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Unions and the Employment Stability of Nonunion Workers: A Sequential Logit Approach

by

Lawrence M. Kahn and Kimio Morimune*

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Abstract

This paper uses 1967 Survey of Economic Opportunity data to estimate the union effect on nonunion employment stability, approximated by a worker's number of spells of unemployment in 1966. A sequence of logit functions is used to estimate the conditional probability of continuing to each successive spell of unemployment. Unions are found to substantially increase the spells of unemployment for nonunion nonwhite workers; effects on whites are found to be weak.

I. Introduction

By causing a union-nonunion wage differential, unions can indirectly lead to frequent employment instability of nonunion workers, as well as influence the latter's wages.[1] This possible effect is of importance for at least two reasons. First, it has been argued that racial differences in unemployment rates are primarily due to differences in frequency rather than duration of unemployment (Flanagan, 1973).[2] A study of union impact on frequency of nonunion unemployment can shed light on this determinant of racial differences. Second, firms are likely to judge job applicants at least partly on the basis of their past employment records; therefore, unions can, through their effect on job stability of nonunion workers, influence the latter's probability of being hired in desirable jobs. Further, opportunities for acquiring firm-specific training depend on stability of employment. Thus, unions can influence the future labor market success of nonunion workers, in addition to current nonunion wages.

This paper uses 1967 Survey of Economic Opportunity (SEO) and labor market data to estimate the impact of unions on the employment stability of nonunion workers. In the SEO, this variable is approximated by the number of spells (S) of unemployment in 1966 that an individual experienced. Since S is a discrete variable taking on one of four possible values, estimation of its determinants requires a limited dependent variable technique. A sequential logistic model is therefore employed.

II. Union and Nonunion Employment Stability

This section discusses a mechanism whereby unions increase nonunion job instability. In addition, possible race-sex differences in

the union impact on nonunion employment stability are explored.

The theoretical framework for analyzing the union effect on nonunion employment stability is based on Mincer's (1976) model of the unemployment effects of minimum wage laws. Suppose that before unions become organized, all workers earn a wage of W_o. Suppose unions then organize a part of the economy and raise union wages to $W_u > W_o$, while nonunion wages become $W_n < W_u$ (W_n may be less than or greater than or equal to W_o--Kahn, 1976).[3] The quantity of union labor demanded will decline. If workers are risk-neutral, then the union-nonunion wage differential will cause frictional unemployment as nonunion workers "wait" for jobs in the union sector. For suppose the probability of finding a nonunion job is 1, while that for a union job is p. Then risk neutrality implies:

$$pW_u = W_n$$

Suppose E_u is total employment, δ is the separation rate in the union sector and that vacancies are filled immediately. Then, following Mincer (1976), if there are still U unemployed job searchers after the vacancies are filled, then,

$$1) \quad p = \frac{\delta E_u}{U + \delta E_u} .$$

If $\dot{W} = \frac{W_u - W_n}{W_n}$, k = fraction of employment that is unionized, and $U_a = \frac{\text{total economy unemployment}}{\text{total economy employment}}$, then it is easily shown that

$$2) \quad U_a = k\delta\dot{W} \text{ (Mincer, 1976).}$$

Thus U_a is the volume of unemployment relative to employment that is caused by the presence of unions compared to the situation where k = 0.

A related measure of union impact is to assess $\frac{dU_a}{dk}$, the effect on U_a of an exogenous increase in k from its present (positive) value.[4] We therefore have:

$$3) \quad \frac{dU_a}{dk} = \delta \dot{W} + k\delta \frac{d\dot{W}}{dk} > 0$$

since both terms on the right hand side of 3) are positive.[5] Thus an increase in the degree of unionism in the economy leads to increasing search unemployment among nonunion workers.[6]

Consideration of those industries that are most likely to increase their degrees of unionism when labor laws become more pro-union reinforces the basic result. Since unions are most likely to organize stable industries, an increase in k implies that the "typical" nonunion employee is more likely to be in an unstable job and to experience involuntary unemployment.[7]

Equation 3) predicts a positive causal relationship from increased union coverage to increased unemployment. Since unemployed workers do not have unlimited resources and since capital markets are imperfect, job searchers may be forced to temporarily take low-paying jobs before finding a more desirable one.[8] Therefore, increases in k should lead to both increasing frictional unemployment and to larger numbers of spells of unemployment for nonunion workers.

Unions can differentially affect the employment stability of nonunion workers in different race-sex groups through at least two mechanisms. First, the union contribution to the level of unemployment may differ by race-sex groups. Second, given the level of unemployment, the expected incidence may differ across groups. These two possibilities are discussed in turn. First, an increase in k can affect U_a differentially if the race-sex groups perceive different returns to

searching. This outcome may occur if the anticipated probability of finding a union job and/or the perceived wage increase from finding one differs by labor market subgroup. In order to predict the effects of unionism on U_a by race-sex group, one must therefore devise a theory to explain the formation of these perceptions. As an approximation to such a theory, we will examine two polar cases:

i) Each race-sex group reacts only to outcomes facing that group (k, δ, $\dot{W}$, $\frac{d\dot{W}}{dk}$).[9]

ii) Each race-sex group reacts only to outcomes for the entire labor market.[10]

In reality, workers in a given group may react <u>both</u> to changes in conditions facing that group and in those facing other groups. However, some interesting findings result from the consideration of the two polar cases. In particular, in case i), we must examine k, δ, $\dot{W}$ and $\frac{d\dot{W}}{dk}$ for each race-sex group.

Ashenfelter (1972) has found that k is roughly 31-32% for black and for while males and 12-13% for black and for white females. It has been suggested that separation rates are larger for females than males and blacks than whites (Barnes and Jones, 1974; Flanagan, 1973). Finally, Ashenfelter (1972) has found the following ranking of $\dot{W}$ in descending order: black males, white males, white females, black females. The authors know of no evidence on $\frac{d\dot{W}}{dk}$ by race-sex group. These findings suggest that U_a and $\frac{dU_a}{dk}$ should be higher for nonwhite than white males. The theoretical results for the other groups in case i) are ambiguous because the rankings of k, δ and $\dot{W}$ among them differ. In case ii), workers perceive neither different probabilities of finding a union job nor different union wages. Thus, only the relationships

among average union wages (for the labor market), the subgroups' average nonunion wages, and k can explain differences in U_a and $\frac{dU_a}{dk}$. In the SEO sample, average nonunion wages can be ranked by group in descending order: white males, nonwhite males, white females, and nonwhite females. Thus, case ii) implies that U_a and $\frac{dU_a}{dk}$ should be less for white males than the other groups. Consideration of these two cases leads to the following tentative conclusion. The impact of unions on overall nonunion unemployment should be stronger for nonwhite males than for white males.

The second stage of the analysis of the union impact on nonunion spells of unemployment by race-sex group involves differences in the expected number of spells given the level of unemployment. A given level of U_a probably results in more spells of unemployment for nonwhites than whites because the liquidity constraint is more likely to be binding for the former group. Because of this consideration and in light of the analysis in cases i) and ii) above, the following hypothesis is suggested. Unionism is most likely to increase nonwhite male spells of unemployment by more than those of white males and to increase nonwhite female spells by more than those of white females.

III. Framework for Estimation

Because stability of employment is approximated by a discrete variable, a qualitative dependent variable approach is suggested. Spells of unemployment, S, takes on four values: 0, 1, 2, and 3 (the last value refers to "3 or more" spells). In order to analyze the determinants of S, the following sequential process is assumed. By

definition, individuals begin each year with 0 spells of unemployment for that year. Throughout the course of the year, some workers experience spells of unemployment, while others don't. Of those who experience unemployment (i.e. $S > 0$), some have only one spell, while others become unemployed more than once ($S > 1$). Finally, of those who have become unemployed twice, some experience additional spells ($S = 3$), while some do not ($S = 2$). Let X be a vector of M explanatory variables. Amemiya (1975) implies that such a process can be analyzed by estimating the following sequence of conditional probabilities:

$$4) \quad q_0(X) = p(S = 0 | X)$$

$$5) \quad q_1(X) = p(S = 1 | S \geq 1, X)$$

$$6) \quad q_2(X) = p(S = 2 | S \geq 2, X).$$

Sequential probabilities q_0, q_1 and q_2 are analogous to transition probabilities in Markov models. However, certain transitions are ruled out (e.g., the transition from 2 to 1 spell of unemployment). Using the calculated results for q_0, q_1 and q_2 the impact of unionism on nonunion spells of unemployment can be estimated as follows. Let $P_i(X)$ be the probability that $S = i$, $i = 0, \ldots, 3$, for nonunion workers, given X. Then

$$7) \quad E(S|X) = \sum_{i=0}^{3} i \, P_i(X)$$

Suppose that X_u is an element of X. Then,

$$8) \quad \frac{\partial E(S|X)}{\partial X_u} = \sum_{i=0}^{3} i \, \frac{\partial P_i(X)}{\partial X_u}$$

is the effect of X_u on nonunion workers' employment instability. By estimating the model of sequential probabilities ($q_0(X)$, $q_1(X)$, $q_2(X)$), the $P_i(X)$ can be estimated, since:

9) $P_0(X) = q_0(X)$

10) $P_1(X) = q_1(X)(1-q_0(X))$

11) $P_2(X) = q_2(X)(1-q_0(X))(1-q_1(X))$

12) $P_3(X) = (1-q_2(X))(1-q_1(X))(1-q_0(X))$.

To estimate the partial derivatives $\frac{\partial E(S|X)}{\partial X_u}$, u = 1, ..., M, a two-step procedure is needed. First, the sequential probabilities $q_i(X)$ are estimated. The logistic distribution is assumed in estimating each $q_i(X)$. Thus:

13) $q_i(X) = \frac{e^{\sum_u \alpha_{iu} X_u}}{1+e^{\sum_u \alpha_{iu} X_u}}$, u = 1, ..., M and i = 0, 1, 2.

Second, the maximum likelihood estimates a_{iu} of the α_{iu} are used in conjunction with equations 8-13 to estimate the $\frac{\partial E(S|X)}{\partial X_u}$. Let $\hat{s}$, $\hat{q}_i$, and $\hat{P}_n$ be respectively the predicted values of s, q_i, and P_n (n=0,...,3) associated with the a_{iu} and $\bar{X}$, the mean of X. Then

14) $\frac{\partial E(S|\bar{X})}{\partial X_u} = -\hat{P}_3\, a_{2u}\, \hat{q}_2 - (\hat{P}_2 + 2\hat{P}_3)\, a_{1u}\, \hat{q}_1 - (\hat{P}_1 + 2\hat{P}_2 + 3\hat{P}_3)\, a_{0u}\, \hat{q}_0$.

Calculation of the exact asymptotic standard errors for this point estimate would be impossible because equation 17 involves products of functions of regression coefficients which are themselves correlated. Instead, standard errors are approximated via the following Taylor expansion procedure. The right hand side of equation 14 contains terms such as $a_{0u}\, \hat{q}_0$, which can be approximated as:

15) $a_{0u}\, \hat{q}_0 \simeq \alpha_{0u}\, q_0(\bar{X}) + a_{0u}\, q_0(\bar{X})\bar{X}\, (a_0 - \alpha_0) + q_0(\bar{X})\, (a_{0u} - \alpha_{0u})$, where

$\bar{X}$ is the mean row vector of X and α_0 and a_0 are respectively the true and estimated column vectors of coefficients from the equation estimating q_0.[11]

Thus:

16) $(a_{0u}\hat{q}_0 - \alpha_{0u} q_0(\bar{X})) \simeq [\alpha_{0u} q_0(\bar{X})(1-q_0(\bar{X}))\bar{X} + (0, \ldots, 0, q_0(\bar{X}), \ldots, 0](a_0-\alpha_0)$

where $q_0(X)$ in the second term in brackets is placed in the same position as a_{0u} is in a_0. Hence,

17) $a_{0u}\hat{q}_0 - \alpha_{0u} q_0(\bar{X}) \simeq C'(a_0 - \alpha_0)$, where C' is the row vector in brackets on the right hand side of equation 16. An approximation to the asymptotic variance of $a_{0u}\hat{q}_0$, then, is $C'[E((a_0 - \alpha_0)(a_0 - \alpha_0)')]C$, where the term in brackets is the asymptotic variance-covariance matrix of a_0, and α_0 and $q_0(\bar{X})$ are replaced in C by a_0 and $\hat{q}_0$, respectively. If one assumes that the regression coefficients across equations are independent, then the approximate asymptotic variance of $\frac{\partial E(\hat{S}|X)}{\partial X_u}$ can be calculated through repeated use of the Taylor expansion procedure.[12] It can thus be shown that

18) Asym. Var $\frac{\partial E(S|X)}{\partial X_u} \simeq d_0'[E((a_0-\alpha_0)(a_0-\alpha_0)')]d_0$

$+ d_1'[E((a_1-\alpha_1)(a_1-\alpha_1)')]d_1 + d_2'[E((a_2-\alpha_2)(a_2-\alpha_2)')]d_2$,

d_0, d_1 and d_2 are column vectors and the terms in brackets are the respective asymptotic variance-covariance matrices of a_0, a_1, and a_2.[13]

The sequential logistic model is used (instead of, for example, the multinomial model) for the following reasons. First, because separate functions are used to estimate q_0, q_1 and q_2, the sequential model allows for "state dependence" in estimating the probability of staying at a given number of spells of unemployment. For one's job

search and/or structure of job offers are likely to change after each spell of unemployment.[14] If workers at the beginning of each year planned their spells of unemployment for the year, then the multinomial logit approach would seem appropriate. However, the sequential approach recognizes that workers do not have perfect foresight and can change their plans. A second advantage of the sequential approach is computational simplicity: in sequential models,"the likelihood function [referring to all four probabilities P_0-P_3] can be maximized by maximizing the likelihood of the dichotomous case repeatedly" (Amemiya, 1975, p. 366).[15] Third, although q_0, q_1 and q_2 in the sequential model are assumed to be mutually independent, equations 9-12 show that P_0-P_3 are not independent.[16] Thus the sequential model appears to be no more restrictive in its assumptions about the dependence of P_0-P_3 as the multinomial model.

IV. Data

The vector X consists of observations on the following explanatory variables for individuals in the twelve standard metropolitan statistical areas (SMSA's) covered by the SEO.[17]

CONSTANT = 1 for all respondents

ED = number of grades of school completed

HLT = 1 if health does not limit the kind or amount of work one can do

= 0 otherwise

MARRY = 1 if respondent is married with spouse present

= 0 otherwise

KIDS = number of children

EXP = age - ED - 6 (potential experience)

$(EXP)^2$ = potential experience squared

MIG = 1 if one had moved more than 50 miles in the previous five years

= 0 otherwise

SOUTH = 1 if one lives in the south

= 0 otherwise

SMSAUNION = percentage of private wage and salary workers in one's SMSA who belong to unions (in percentage points for 1967)[18]

SMSAUNEM = 1966 unemployment rate for the SMSA where one lives (in percentage points)

SMSAMANU = 1966 percentage of nonagricultural employment in the respondent's SMSA that is in manufacturing industries (in percentage points)

SMSARACE = 1966 percentage of population in one's SMSA that is nonwhite

Maximum likelihood estimates of the α_{ij} are obtained for each of four race-sex groups: nonwhite males, nonwhite females, white males and white females.

ED, HLT, EXP, MARRY and KIDS all represent individual productivity characteristics and/or constraints on one's labor supply. MIG is included because the union crowding effect may cause out-migration from highly unionized areas. On the other hand, high union wages may cause in-migration to highly unionized areas (Hall, 1970). In either case, the supply of labor in less-unionized labor market areas is affected by actions of unions in highly organized areas. Our control for this phenomenon is similar to the control for migration status in cross-sectional studies of labor force participation rates (Cain, 1967).

One might suppose that the local unemployment rate SMSAUNEM would take this factor into account; however, due for example to information problems, a recent migrant might change jobs more often before finding an "appropriate" one than an otherwise identical permanent resident (Flanagan, 1974). Therefore, we have controlled for migration status.[19]

SOUTH, SMSARACE, SMSANUNEM and SMSAMANU are included to control for local labor market conditions besides unionism. Other things equal, higher SMSARACE is expected to cause less stable nonwhite employment for two reasons. First, SMSARACE approximates the degree of crowding of minorities into low-paying, nonunion jobs that is not caused by unions (Bergmann, 1971). Second, the larger SMSARACE is, other things equal, the more likely are nonwhites to find jobs with discriminating rather than nondiscriminating employers (Becker, 1957). For either of these reasons, one would expect high levels of SMSARACE to contribute (cereris paribus) to more rapid turnover for nonwhite workers.[20] If racial employment segregation is pervasive (or was in 1966, the year for which our data on spells of unemployment were collected), then white workers in areas with high values of SMSARACE face less competition than otherwise. Therefore, one might expect white spells of unemployment to be negatively related to SMSARACE (other things equal). On the other hand, whites in areas with high values of SMSARACE may have better opportunities than otherwise and quit more often. Further, if employers in the nonunion sector view whites and nonwhites in that sector as substitutable, then a higher value of SMSARACE would lead, to higher white as well as nonwhite S. Therefore the impact of SMSARACE for whites is ambiguous.

Other things equal, higher values of SMSAUNEM should lead to higher nonunion S (or lower labor force participation rates) due to competition for jobs. If manufacturing is a high-wage sector, then nonunion workers in SMSA's with large values of SMSAMANU may very well quit more often than otherwise. This outcome is similar to the one in Mincer's (1976) minimum wage model.

Finally, the coefficients for SMSAUNION represent the effect of a union environment on the job instability of nonunion workers. For reasons stated above, the coefficient for this variable is expected to be larger for nonwhite than for white men and larger for nonwhite than for white women. From the earlier discussion, one might expect the spells of any race-sex group to depend on the degree of unionism of each race-sex group. In principle, then, in the estimation for each subgroup one might use the four values of SMSAUNION as explanatory variables. However, because the SEO sample contains only twelve SMSA's, and because the rankings of SMSAUNION by race-sex group across SMSA's are extremely consistent, such a procedure is not feasible.[21]

Because SMSAUNION and SMSAUNEM are both in principle endogenous to our model, possible biases arise in the estimation of the effect of unions on nonunion employment stability. First, if unions are better able to organize stable than unstable workers, then SMSA's with high values of S are less likely to become highly unionized.[22] In this case, a negative simultaneity bias is imparted to the effect of unions on nonunion employment instability. Since industrial composition is probably a major determinant of SMSAUNION, the control for SMSAMANU at

least partially corrects for factors that influence the degree of unionism. It is possible that high wages in an SMSA lead to a high degree of unionism and to search unemployment, producing a positive bias on SMSAUNION (Ashenfelter and Johnson, 1972). However, the strength of such a bias is at least in part discounted by Kahn's (1976) result that high degrees of unionism, ceteris paribus, led to lower real wages for nonwhite male and white female nonunionists.

Second, the unemployment rate of an area is in part determined by the crowding effect of unions.[23] Inclusion of SMSAUNEM, then, appears to be another source of downward bias on the SMSAUNION coefficient because part of unionism's effect on the S of nonunion workers operates through unemployment. SMSAUNEM is included to account for exogenous factors (such as age composition of the labor force and detailed industrial composition) that influence an area's unemployment rate. Since the sources of possible bias mentioned here are largely negative, our confidence in a positive union effect on S would be strengthened.[24]

V. Empirical Results

Appendix 3 summarizes the results by race-sex group for the sequential probabilities q_0, q_1 and q_2.[25] Table 1 presents estimates and approximate asymptotic standard errors for the impact of each explanatory variable on spells of unemployment.[26]

The coefficients for SMSAUNION in Table 1 measure the estimated impact on S of a one percentage point change in SMSAUNION. In the SEO sample, this variable ranges from 10% to 46%. A conservative estimate, then, of the difference in unionism between a highly unionized SMSA

Table 1[c]: Estimates of $\frac{\partial(E(\hat{S}|X))}{\partial X_u}$ by Race-Sex Group

Independent variable	Race-Sex Group			
	Nonwhite males	Nonwhite females	White males	White females
ED	-.029 (.009)	-.012 (.009)	-.021 (.003)	-.013 (.004)
EXP	.001 (.007)	-.013 (.006)	-.009 (.002)	-.007 (.003)
$(EXP)^2$	-.0001 (.0002)	.0002 (.0001)	.0002 (.0001)	.00009 (.00005)
MIG	.061 (.065)	.002 (.054)	.068 (.022)	.019 (.024)
HLT	-.062 (.097)	.003 (.069)	-.030 (.034)	-.034 (.027)
SMSAUNEM	.154 (.072)	-.012 (.048)	.005 (.020)	.010 (.023)
SMSAUNION	.009 (.0059)	.0073 (.0038)	.0009 (.002)	-.0004 (.002)
SMSARACE	.023 (.015)	-.018 (.010)	-.003 (.004)	-.009 (.006)
SMSAMANU	-.003 (.005)	-.0009 (.003)	-.0002 (.002)	-.003 (.002)
SOUTH	.267 (.139)	.217 (.101)	-.066 (.080)	.094 (.071)
KIDS	--	-.004 (.011)	--	.006 (.007)

[c]Approximate asymptotic standard errors in parentheses. Mean values of the independent variables were used in the calculations.

and a relatively non-unionized SMSA would be 20 percentage points. Table 2 illustrates the impact on spells of unemployment of such a difference in unionism.

Table 2: Percentage Impact of a 20 Percentage Point Change in SMSAUNION on $E(\hat{S}|\overline{X})$ (Ceteris paribus)

Race-Sex Group	Union Effect
Nonwhite males	63.0%
Nonwhite females	69.4%
White males	17.3%
White females	-8.5%

Tables 1 and 2 suggest that the degree of unionism strongly increases the expected spells of unemployment for nonwhite males and females. For these groups the point estimates are large and sizable compared to their standard errors. These results are in principle strengthened by the discussion of downward biases in Section IV above.[27]

In assessing the impact of unionism on racial differentials in nonunion spells of unemployment one can pose the following questions.[28] First, if nonwhites' and whites' spells of unemployment were equally affected by unionism, what impact would there be on these differentials? Such a comparison is analogous to the coefficient effect in decomposing racial wage differentials (Flanagan, 1974). Second, if nonwhites had both the white coefficients for SMSAUNION and the white mean SMSAUNION, what would racial differentials in spells be?[29]

Table 3 illustrates the overall differences by race and sex in average spells of unemployment.

Table 3*: Average Number of Spells of Unemployment for Nonunion Workers, 1966

Race-Sex Group	Mean Spells	Predicted Spells**
Nonwhite males	.286	.290
Nonwhite females	.259	.209
White males	.104	.102
White females	.094	.089

*Source: 1967 SEO Tape.

**All explanatory variables are assumed to equal their mean values.

Table 4 presents the hypothetical nonwhite spells of unemployment under the assumptions of the two questions posed above.

Table 4: Hypothetical Nonwhite Spells of Unemployment

Group	$E(\hat{S}\|\overline{X})$ with White SMSAUNION Coefficient	$E(\hat{S}\|\overline{X})$ with White Mean and Coefficient for SMSAUNION
Nonwhite males	.160	.141
Nonwhite females	.157	.159

Table 4 illustrates that if nonwhites responded to unionism in the same way as whites, then the race differentials in spells would be substantially reduced.[30] The mechanism by which the differentials would be reduced is not entirely clear: unionism leads to spells of unemployment via k, δ, $\dot{W}$, $\frac{d\dot{W}}{dk}$ and liquidity considerations. However,

one can conclude that if these factors were equal for whites and nonwhites within sex groups, then race differentials in the incidence of unemployment would be greatly reduced.

VI. Conclusions

The major result obtained here has been the strong negative effect of unions on the employment stability of nonunion nonwhite workers. This finding is consistent with Hall's (1970) view of high rates of nonwhite unemployment at 'full employment':

> The true problem of hard-core unemployment is that certain members of the labor force account for a disproportionate share of unemployment because they drift from one unsatisfactory job to another, spending the time between jobs either unemployed or out of the labor force (Hall, 1970, p. 389).

Further, this paper has employed a technique that has many potential applications to situations involving sequential choice.

Footnotes

1. Unions can in principle have opposing effects on the wages and working conditions of nonunion workers. On the one hand, if unions raise their members' wages, less labor is hired in the union sector than would otherwise be the case. Workers may then crowd into the nonunion sector, leading to lower wages and worse working conditions there (the "crowding effect"). On the other hand, union gains may compel employers of nonunion labor to raise wages and improve working conditions in order to keep unions out or to reduce turnover (the "threat effect"). At any time, each of these two effects may coexist in the labor market. In an earlier paper, the net effect of unions on the wages of nonunion workers in relatively unorganized sectors was investigated (Kahn, 1976). It was found that unions lowered the real wages of nonunion white females and nonwhite males and raised those of white males.

2. In a steady state sense, the unemployment rate equals frequency times duration.

3. In reality, unions are more likely to arise in high wage industries (see Ashenfelter and Johnson, 1972 or Kahn, forthcoming). By assuming that preunionization wages were W_o for all workers, this model assesses the impact of a union-caused wage differential on nonunion employment stability. Most single equation studies of union impact indeed found that unions raised their members' relative wages (see Ashenfelter and Johnson, 1972). When unionism is viewed as endogenous, the results are mixed. In the one hand, Ashenfelter and Johnson (1972) and Schmidt and Strauss (1976) could not reject the hypothesis that unions do not raise relative wages. On the other hand, when changes in technique of production and labor quality were allowed, Kahn (forthcoming) was able to reject the null hypothesis of no union impact.

4. Caused, say, by a change in labor laws.

5. Hendricks (1975) indeed found that union wages in an industry were positively affected by the degree of unionism in the industry and in the labor market as a whole. Rosen (1969) and Stafford (1968) found that the union-nonunion differential in an industry rose as the degree of unionism in the industry rose.

6. It can be shown that the qualitative result stated here holds in the cases of fixed or variable labor force size (Miner, 1976).

7. Ashenfelter and Johnson (1972) found that the nonunion sector is more competitive than the union sector. Thus higher firm mortality rates imply that nonunion workers are more likely to lose their jobs.

8. See Lippman and McCall (1976) for a discussion of job search models in which the possibility of bankruptcy forces a searcher's reservation wage downward as search time increases.

9. Such would be the case if race-sex groups perceive the labor market to be segmented by race and/or sex.

10. In this case, for example, a rise in $\dot{W}$ for nonwhite males affects all the subgroups equally.

11. To see that expression 15 holds, consider the following expansion around α_o:

$$\hat{a_{0u}q_0} \tilde{} \alpha_{0u}q_0 + \frac{\partial(\alpha_{0u}q_0)}{\partial\alpha_0}\Big|_{\alpha_0} Q_0\ (a_0-\alpha_0) = \alpha_{0u}q_0 + q_0(a_{0u}-\alpha_{0u})$$

$$+ \frac{\partial(\alpha_{0u}q_0)}{\partial\bar{X}\alpha_0} \cdot \frac{\partial(\bar{X}\alpha_0)}{\partial\alpha_0}\Big|_{\alpha_0}\ \alpha_0\ (a_0-\alpha_0) = \alpha_{0u}q_0 + q_0(a_{0u}-\alpha_{0u})$$

$+ q_0(1-q_0)\bar{X}(a_0\alpha_0)$, where q_0 is evaluated at $\bar{X}$.

12. The assumption of the independence of a_0, a_1 and a_2 is discussed below.

13. See Appendix 1 for a derivation of expression 18.

14. Heckman and Willis (1977), in a model of female labor force participation, suggest that heterogeneity of the sample (not accounted for by the explanatory variables) can cause an apparent state dependence of transition probabilities, even though the underlying model predicts state independence. It is argued here that even for identical individuals, the probability of say, moving from one to two spells of unemployment may differ from the probability of moving from two to three spells.

15. One might suppose that the following ordered response model would apply to our case (this discussion follows that of Amemiya, 1975). Suppose one can be unemployed 0, 1 or 2 times and that the expected number of spells is a function of a single variable z. Individual i becomes unemployed twice if $u_t < z_t$ and at least once if $u_t - c < z_t$, for u_t, a random variable and c, a positive constant. If u_t is distributed as logistic, then we have:

$P\ (S=2) = L\ (z)$

$P\ (S=1 \text{ or } 2) = L\ (z+c)$

$P\ (S=0) = L\ (-z-c)$, where L (-) is the logistic distribution function. This ordered response model is not used here because

the assumption of the constant shift parameter, c, is too restrictive. In equations 9-12 above, no particular restrictions are placed, for example, on the difference between P(S=1) and P(S≥1). In addition, according to Amemiya,

> If one uses an ordered model when the true model is unordered, one could get into serious trouble, whereas the loss in using an unordered model in the ordered situation is only in efficiency since consistent estimates are still obtained (Amemiya, 1975, p. 366).

16. In principle, the errors in the sequential probability functions are correlated across equations. For example, the random factors (personality, etc.) that make one unlikely to experience zero spells of unemployment may also decrease the probability of having only one spell. However, the three probability functions corresponding to q_0, q_1 and q_2 are not all estimated from the same sample. There is, therefore, no obvious way within the sequential logistic model of taking into account the across-equation error structure.

17. See Appendix 2 for derivations and sample mean values for X.

18. Equation 3) above implies that k^2 should be an explanatory variable (because $\frac{dU_a}{dk}$ is a function of k). However, equation 3) refers to overall unemployment, not to spells of unemployment. The functional form of the relationship between spells and k is not intuitively obvious. Therefore, for reasons of computational ease, we have not included an SMSAUNION squared term in X.

19. Migrants may also have different productivity characteristics from nonmigrants (e.g., the former may be more aspiring). The MIG coefficients do not measure the impact of migrating; instead they compare the respective probabilities of "stopping" at 0, 1 or 2 spells of unemployment (ceteris paribus) of recent migrants to those of nonmigrants in the SMSA to which the former have migrated.

20. If such turnover results in decreased labor force participation rates, then S (the number of spells of unemployment) under-states the impact of SMSARACE on employment stability.

21. The controls for SMSAUNEM, SMSARACE and SMSAMANU leave us with essentially eight degrees of freedom. SMSAUNION for the whole market is used, rather than that for the particular subgroup being considered, because this is the more general specification. That is, for example, the unionism of each race-sex group affects nonwhite males by influencing the total

SMSAUNION. In addition, since it was argued above that illiquidity led to higher S, it would have been interesting to use family income and assets as explanatory variables. However, these were not available on the SEO tape we used.

22. In the long run, the "typical" nonunion worker may have a greater or smaller propensity to be stable than the average union worker in the same SMSA. On the one hand, unions are ceteris paribus more likely to recruit in industries with stable employees, and union-won benefits may increase that stability. On the other hand, Mincer's model implies that the least risk averse workers are most likely to apply to union firms for jobs. These employees may well be more unstable than more risk averse workers. Thus, there is little reason to believe that unions will lower the average propensity of nonunion workers to be stable. Instead, it is argued here that unions' effect on the environment of nonunion workers leads to increased stability.

23. Industrial composition also affects an area's unemployment rate: for example, areas with seasonal industries should, other things equal, have higher unemployment rates than otherwise.

24. Notice that intervening variables such as average union or nonunion wages are not included as explanatory variables. This is the case because one is interested in the ultimate impact of personal characteristics and labor market variables on nonunion employment stability. In particular, SMSAUNION may directly and/or indirectly (through union wages, nonunion wages, etc.) influence the dependent variable. This formulation allows estimation of the direct plus indirect effects of unionism on spells of unemployment. In addition, the assignment of S=3 to those with more than three spells of unemployment further understates the union impact on spells. It is more likely, other things equal, that nonunion workers with S=3 in highly unionized SMSA's have more than three spells than those with S=3 in nonunionized SMSA's.

25. Because of the nature of the computer program used to estimate the q_i, the predicted probabilities $\hat{q}_i$ are computed as follows:

$$\hat{q}_i = \frac{e^{2\Sigma_j a_{ij}\bar{X}_j}}{1+e^{2\Sigma_j a_{ij}\bar{X}_j}} \qquad i = 0,1,2 \text{ and } j = 1, \ldots, 13,$$

where the a_{ij} are the reported maximum likelihood estimates.

26. In Table A.3.1. in the estimation of q_2 for white males and white females, there are some empty cells (SOUTH for white males and SOUTH and HLT for white females). In the calculations of Table 1, it is assumed that the coefficients for these variables are 0 and that the covariances between these coefficients and others are 0. This problem could have been avoided

for these two groups by restricing S to the values 0, 1 and 2 (the last value refers to "2 or more" spells). When this procedure is used, the results are qualitatively unchanged.

27. The standard errors on the local labor market variable coefficients (including SMSAUNION) might well have been smaller if the SEO sample had included more than 12 SMSA's.

28. We concentrate on race differences because, as will be seen, sex differences in spells are not large and because the functional forms used here differ by sex. The data do not permit an estimate of the hypothetical differentials in spells when SMSAUNION = 0 because this value is outside of the range for our sample.

29. White coefficients and means are used as benchmarks because whites are a majority of the labor force.

30. The effect of the white mean is to further lower the predicted spells of nonwhite males but not to noticeably affect that of nonwhite females.

Appendix 1: Approximation of the Asymptotic Variance of $\frac{\partial \hat{E}(S|\bar{X})}{\partial X_u}$

Equation 14 gives an expression for $\frac{\partial \hat{E}(S|\bar{X})}{\partial X_u}$. Using expressions 14 and 15, one can write the following approximation:

A1.1): $\frac{\partial \hat{E}(S|\bar{X})}{\partial X_u} - \frac{\partial E(S|\bar{X})}{\partial X_u} \approx$

$$\hat{P}_3 \hat{q}_2 [a_{0u} \hat{q}_0 + a_{1u} \hat{q}_1 + a_{2u} (2\hat{q}_2 - 1)] \bar{X} (a_2 - \alpha_2)$$

$$+ \hat{q}_1 [(\hat{P}_2 + 2\hat{P}_3) \hat{q}_o a_{0u} + (\hat{P}_2 + 2\hat{P}_3)(2\hat{q}_1 - 1) a_{1u} + \hat{P}_3 \hat{q}_2 a_{2u}] \bar{X} (a_1 - \alpha_1)$$

$$+ \hat{q}_0 [(\hat{P}_1 + 2\hat{P}_2 + 3\hat{P}_3)(2\hat{q}_0 - 1) a_{0u} + (\hat{P}_2 + 2\hat{P}_3) \hat{q}_1 a_{1u} + \hat{P}_3 \hat{q}_2 a_{2u}] \bar{X} (a_0 - \alpha_0)$$

$$- \hat{P}_3 \hat{q}_2 (a_{2u} - \alpha_{2u})$$

$$- (\hat{P}_2 + 2\hat{P}_3) \hat{q}_1 (a_{1u} - \alpha_{1u})$$

$$- (\hat{P}_1 + 2\hat{P}_2 + 3\hat{P}_3) \hat{q}_0 (a_{0u} - \alpha_{0u}).$$

Under the assumption of independence across equations of the logit coefficients, expression A1.1 allows one to approximate the asymptotic variance of $\frac{\partial \hat{E}(S|\bar{X})}{\partial X_u}$, as summarized by expression 18.

Appendix 2: Description of the Data

This appendix describes the sources and derivations of several of the variables used in this study.

EXP: Potential experience is constrained to be nonnegative and to begin at age 16.

SOUTH: In this study, Houston is taken to be in the South, but Baltimore and Washington, D.C. (considered Southern SMSA's by the Census) are not. One usually controls for Southern residence to take into account the following factors: cost of living, rural residence, industrial structure, degree of unionism and degree of racism. Controlling for the 2nd, 3rd, and 4th factor (as this study does), we believe that Baltimore and Washington, D.C. are more similar in labor market structure to the other SMSA's of the SEO than these two SMSA's are to Houston. Cost of living is not directly relevant to this study.

SMSAUNION: This variable was calculated from the SEO.

SMSAUNEM: This variable was taken from Hall (1970).

SMSAMANU: This variable was calculated from USBLS (1966).

SMSARACE: This variable was constructed via an interpolation of data from 1960 (U.S. Bureau of the Census, 1960) and 1970 (U.S. Bureau of the Census, 1970). Total and nonwhite population were assumed to have grown at constant rates from 1960 to 1970.

Mean Values of Independent Variables

Variable	White females	Nonwhite females	White males	Nonwhite males
ED	11.6637	10.1634	12.1288	9.76583
EXP	18.9576	21.4630	18.3826	19.0299
$(EXP)^2$	562.100	637.282	575.658	534.194
MIG	.169462	.129377	.214519	.189964
HLT	.940865	.89786	.93002	.929510
SMSAUNEM	3.71253	3.34805	3.6416	3.29785
SMSAUNION	31.0159	28.2646	30.2479	26.9391
SMSARACE	15.5628	17.8873	15.7494	18.1072
SMSAMANU	27.9537	26.2482	28.4519	25.3978
SOUTH	.0317741	.0875486	.0477436	.111111
KIDS	1.26567	1.86673	---	---

Appendix 3: Empirical Results for the Sequential Probabilities

Table A.3.1. Logit Results for Sequential Probabilities[a]
Nonwhite Males

Independent Variables	Sequential Probabilities		
	q_0	q_1	q_2
Constant	0.8036 (1.0433)	3.9489 (2.2784)	6.6531 (4.5140)
ED	0.0503 (0.0162)	0.0376 (0.0382)	-0.0020 (0.0575)
EXP	0.0081 (0.0128)	-0.0468 (0.0261)	0.0306 (0.0437)
$(EXP)^2$	0.0002 (0.0003)	0.0007 (0.0006)	-0.0009 (0.0010)
MIG	-0.1612 (0.1158)	0.0984 (0.2253)	0.1404 (0.3572)
HLT	0.2628 (0.1706)	-0.2579 (0.3549)	-0.9139 (0.5594)
SMSAUNEM	-0.1097 (0.1214)	-0.4989 (0.2757)	-1.0076 (0.5706)
SMSAUNION	-0.0093 (0.0103)	-0.0298 (0.0210)	-0.0198 (0.0328)
SMSARACE	-0.0162 (0.0263)	-0.0857 (0.0558)	-0.1202 (0.1098)
SMSAMANU	0.0010 (0.0080)	0.0205 (0.0167)	-0.0045 (0.0247)
SOUTH	-0.2321 (0.2432)	-0.9058 (0.5035)	-0.9958 (0.8310)
KIDS[b]	---	---	---
n	837	139	69

[a]Asymptotic Standard errors in parentheses;
n = no. of observations

[b]This variable was unavailable for males.

Table A.3.1. (Continued)
Nonwhite Females

Independent Variables	Sequential Probabilities		
	q_0	q_1	q_2
Constant	0.3360 (0.8949)	-0.9849 (2.112)	-5.449 (4.7061)
ED	0.0225 (0.0203)	0.0236 (0.0509)	0.0632 (0.0790)
EXP	0.0299 (0.0132)	0.0120 (0.0301)	0.0391 (0.0482)
$(EXP)^2$	-0.0002 (0.0003)	-0.0006 (0.0007)	-0.0012 (0.0011)
MIG	-0.0063 (0.1293)	-0.0172 (0.2599)	0.0719 (0.3966)
HLT	0.0122 (0.1629)	-0.0087 (0.3373)	-0.2250 (0.5477)
SMSAUNEM	-0.0013 (0.1062)	0.0499 (0.2556)	0.2792 (0.5226)
SMSAUNION	0.0013 (0.0089)	-0.0684 (0.0193)	-0.0507 (0.0328)
SMSARACE	0.0084 (0.0216)	0.0966 (0.0525)	0.2364 (0.1447)
SMSAMANU	-0.0112 (0.0077)	0.0436 (0.0192)	0.0338 (0.0349)
SOUTH	-0.1159 (0.2375)	-1.5141 (0.5032)	-1.2729 (1.0411)
KIDS	-0.0054 (0.0237)	0.0410 (0.0588)	0.0870 (0.1039)
n	1028	146	63

Table A.3.1. (Continued)
White Males

Independent Variables	Sequential Probabilities		
	q_0	q_1	q_2
Constant	-0.3167 (0.8603)	-0.8752 (1.6407)	-0.7562 (3.1057)
ED	0.0869 (0.0139)	0.0777 (0.0316)	0.1128 (0.0725)
EXP	0.0508 (0.0114)	-0.0239 (0.0246)	0.0796 (0.0567)
$(EXP)^2$	-0.0008 (0.0003)	0.0005 (0.0006)	-0.0022 (0.0015)
MIG	-0.3100 (0.1047)	-0.1573 (0.2129)	-0.1672 (0.3800)
HLT	0.2681 (0.1593)	-0.1909 (0.3218)	-1.1312 (0.6027)
SMSAUNEM	-0.0793 (0.0924)	0.1444 (0.1785)	0.3514 (0.3304)
SMSAUNION	-0.0063 (0.0089)	0.0118 (0.0183)	-0.0233 (0.0282)
SMSARACE	0.0164 (0.0203)	-0.0000 (0.0365)	-0.0369 (0.0605)
SMSAMANU	0.0035 (0.0084)	-0.0087 (0.0170)	-.0019 (0.0335)
SOUTH	0.2057 (0.3759)	0.5664 (0.7766)	---[c]
KIDS	---	---	---
n	1529	136	51

[c]In estimating q_2, all white males had SOUTH = 0.

Table A.3.1. (Continued)
White Females

Independent Variables	Sequential Probabilities		
	q_0	q_1	q_2
Constant	0.6627 (0.8347)	-4.1307 (4.4422)	-33.3998 (29.5699)
ED	0.0507 (0.0182)	0.1345 (0.0661)	0.1070 (0.1790)
EXP	0.0464 (0.0124)	-0.0345 (0.0397)	0.0193 (0.1060)
$(EXP)^2$	-0.0007 (0.0003)	0.0008 (0.0009)	0.0001 (0.0022)
MIG	-0.1164 (0.1253)	-0.1939 (0.3473)	0.9321 (0.9387)
HLT	0.1611 (0.1609)	0.2744 (0.5157)	---[d]
SMSAUNEM	-0.1838 (0.0902)	0.2107 (0.3949)	2.5594 (2.4719)
SMSAUNION	-0.0029 (0.0083)	0.0259 (0.0346)	0.0204 (0.0528)
SMSARACE	0.0181 (0.0180)	0.0164 (0.0953)	0.9708 (0.8307)
SMSAMANU	-0.0010 (0.0072)	0.0462 (0.0391)	0.2529 (0.2275)
SOUTH	-0.7141 (0.3689)	0.8649 (1.2248)	---[d]
KIDS	-0.0438 (0.0387)	0.1289 (0.1206)	-0.3318 (0.3321)
n	1133	96	20

[d] In estimating q_2, all white females had HLT = 1 and SOUTH = 0.

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